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Energy prices between Beijing and Shanghai: Testing for the long-run equilibrium

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Abstract

This paper mainly aims to test the long-run relationship between Beijing and Shanghai energy prices. Panel cointegration tests indicate a cointegrating vector. Additionally, feedback was implied for the changes in Beijing and Shanghai energy prices. Therefore, empirical evidence supports the argument that similar macro fundamentals (supreme political power, considerable economic and population sizes) lead to the long-term equilibrium between energy markets. We suggest that the central government plays a leading role in regulating energy prices.

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Keywords: energy price; macro fundamental; long-run relationship; Granger causality; panel cointegration

1. Introduction

Previous studies suggested that housing demand affected energy prices in Shanghai in the long run but not vice versa [1]. New housing supply positively impacted energy prices in Beijing in the long term [2]. These studies were conducted on a specific city basis. However, cities share similar macro fundamentals, including politics, economy and population. Hence, prices between city energy markets significantly impacted each other, which may form a long-run equilibrium. Cointegration argues that similar macro fundamentals can lead to the long-term equilibrium [3]. Two principal cities in China are Beijing and Shanghai. They both hold substantial economic and population sizes, and significant political power. Therefore, this study mainly aims to examine the possible long-run

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equilibrium between changes in energy prices between Beijing and Shanghai energy markets. The study will introduce panel cointegration tests. Also, a bi-directional Granger causality test will be driven.

2. Methodology

We conduct panel unit roots using two techniques [4, 5]. We drive the panel cointegration tests using an Engle-Granger based test, and a Johansen trace based test [6]. The Engle-Granger test is to drive a regression of a time-series variable y_t on another time-series variable x_t and thus, check if the regression residual μ_t is stationary [7]:

$$y_t = \alpha + \beta x_t + \mu_t \tag{1}$$

The Johansen trace test is based on the following error-correction equation [8]

$$\Delta y_t = \mathbf{\Pi} y_{t-1} + \sum_{i=1}^{p-1} \mathbf{\Gamma}_i \Delta y_{t-i} + \mathbf{B} x_t + \varepsilon_t \tag{2}$$

Where $\mathbf{\Pi} = \sum_{i=1}^{p} A_i - I,\ \mathbf{\Gamma}_i = -\sum_{j=i+1}^{p} A_j$, y_t is a k-vector of non-stationary $I(1)$ variables; x_t is a d-vector of deterministic variables, and ε_t is a vector of white noises with zero mean and finite variance. If cointegration exists, we estimate the cointegrating vector using FMOLS [9, 10] and DOLS [11, 12]. According to Engle and Granger [7], cointegration implies at least one directional Granger causality existing between time-series variables. We conduct the Granger causality test [3].

3. Data

The study employed a monthly time series for the period from January 2008 to December 2013. The purchase price of fuels and power represents energy price (*ENERGY PRICE*, the same month of the last year=100). The purchase price of fuels and power was measured as the nominal index change as compared with the same month of last year. The purchase price of fuels and power was contained in industrial producer's purchase price, reflecting variations in the prices of representative energy products that representative companies purchase. These representative energy products include coal, gasoline, and natural gas [13]. Energy prices were deflated by the consumer price index (*CPI*, the same month of the last year=100). Thus, we used real energy prices.

4. Empirical Results

Table 1 reports the results of panel unit root tests. The Levin-Lin-Chu tests suggested two unit roots, but the Breitung tests did not imply any unit root. Hence, we regarded the panel dataset as an $I(1)$ process. The Pedroni tests rejected the null hypothesis of no cointegration. The Fisher tests accepted the hypothesis of one cointegrating vector at the 2% confidence level when we pre-specified a lag of 4 for the test (Table 2). Thus, the series were cointegrated and contained one cointegrating vector. We estimated the cointegrating vector using the FMOLS (Table 3) and DOLS (Table 4) estimators, respectively. Based on the estimates in Table 3 and Table 4, the Granger causality tests strongly rejected the hypothesis of no causality on both directions in all cases (Table 5).

Table 1. The panel unit root tests

	Level			First difference			Second difference		
Method	*k*	test statistic	P-value	*k*	test statistic	P-value	*k*	test statistic	P-value
Levin, Lin and Chu	2	0.12	0.55	2	-0.88	0.19	2	-18.3	0.00
Breitung	2	-2.95	0.00	-	-	-	-	-	-

Notes: Variables were in a logarithmic form (i.e., log(*ENERGY PRICE*)/log(*CPI*)). k is a number of lags. Tests chose k utilizing the modified Schwarz information criterion (MSIC). The equations included both the trend and intercept [14]. P-values are in parentheses.

Table 2. the panel cointegration tests

Lag	Pedroni	Statistic	P-value
11	Panel v-Statistic	3.43	0.00
	Panel rho-Statistic	-3.39	0.00
	Panel PP-Statistic	-3.21	0.00
	Panel ADF-Statistic	-2.98	0.00
	Fisher		
	H_0	Fisher trace statistic	P-value
2	0	37.37	0.00
	⩽1	23.24	0.00
4	0	50.27	0.00
	⩽1	18.42	0.02
6	0	39.99	0.00
	⩽1	26.15	0.00

Notes: For the Pedroni tests, equations included both the trend and intercept. We chose a max lag of 11 using modified AIC (MAIC). For the Fisher combined Johansen test, we used Johansen's Model 4 and thus equations included both the trend and intercept [15, 16]. Tests used a lag of 2, 4 and 6 for the Fisher tests, respectively.

Table 3. Estimates of cointegrating vector utilizing the FMOLS estimator

Variable	Coefficient	Standard error	*t*-statistic	P-value
Dependent: LOG(*BEIJING_ENERGY PRICE*)/LOG(*BEIJING_CPI*)				
LOG(*SHANGHAI_ENERGY PRICE*)/LOG(*SHANGHAI_CPI*)	0.64	0.02	31.46	0.00
Adjusted R-squared	0.86	S.D. dependent var	0.03	
Durbin-Watson stat	0.49	Long-run variance	0.00	

Notes: Test used pooled (weighted) estimation. The test used both the constant and linear trend. Coefficient covariance was computed using homogeneous variance. Long-run covariance estimation selected lag using AIC.

Table 4. Estimates of cointegrating vector utilizing the DOLS estimator

Variable	Coefficient	Standard error	*t*-statistic	P-value
Dependent: LOG(*BEIJING_ENERGY PRICE*)/LOG(*BEIJING_CPI*)				
LOG(*SHANGHAI_ENERGY PRICE*)/LOG(*SHANGHAI_CPI*)	0.70	0.01	60.40	0.00
Adjusted R-squared	0.97	S.D. dependent var	0.03	
Long-run variance	0.00			

Notes: The Test equation contained the linear trend. Test used pooled (weighted) estimation. Coefficient covariance was estimated by homogeneous variance. We chose leads and lags by AIC.

Table 5. The bi-directional Granger causality tests

Hypothesis	Wald- χ^2	P-value
No causality from Shanghai to Beijing*	989.5	0.00
No causality from Beijing to Shanghai*	3753.6	0.00
No causality from Shanghai to Beijing**	3648.2	0.00
No causality from Beijing to Shanghai**	3775.3	0.00

Notes: *Tests were based on estimates in Table 3. **Tests were based on estimates in Table 4.

5. Discussions and Summary

Two cointegration tests strongly supported the existence of a long-run relationship between Beijing and Shanghai energy prices. Estimates using FMOLS and DOLS suggested that a 1% growth in energy prices in Shanghai may lead to 0.64%-0.7% increase in energy prices in Beijing. Additionally, feedback existed between changes in Beijing and Shanghai energy prices.

Therefore, energy prices in Beijing and Shanghai exhibited not only a typical long-run equilibrium but also close short-run dynamics. The empirical evidence supports the argument that Beijing and Shanghai have commonly promoted a long-run growth in energy prices because these two metropolises have similar macro fundamentals: supreme political power in China, and considerable economic and population sizes. A policy implication for this is that fluctuations in energy prices in China's two most important cities may impose a nationwide effect on the energy price across other Chinese cities and thereby the central government should bear the main responsibility to regulate energy prices.

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